# Online Appendix for Free Trade and the Formation of Environmental Policy: Evidence from US Legislative Votes*

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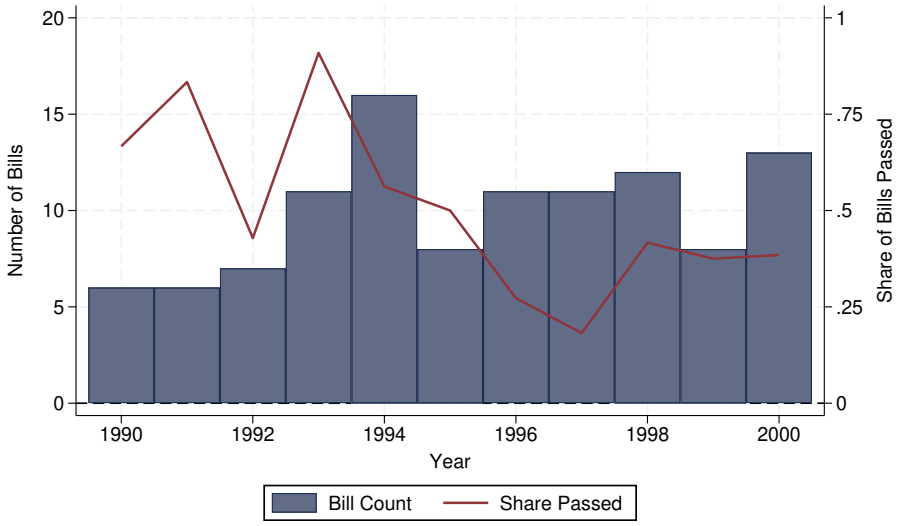
# Appendix A Additional Summary Statistics

This section provides additional summary statistics to complement those presented in the main paper.

First, to highlight how environmental RCVs changed during our period of study, Figure A.1 shows trends in the number of environmental bills that were put forth for an RCV in the House and their outcomes over time. The number of bills put up for an RCV ranges from five to sixteen per year, with an annual average of just under ten and a slight increase over time. Most notable is the stark reduction in the share of bills that passed a simple majority after NAFTA's introduction. In years 1990 to 1993, between 45% and 80% of environmental bills passed a simple majority. Following the implementation of NAFTA, however, this share fell immediately to below 50%, and declined each year until 1997, hovering near 25% for the rest of our sample. Our analysis attempts to determine how much of this change is in fact due to NAFTA by exploiting the geographic variation in exposure to NAFTA across Congressional districts.

Figure A.1: House Environmental Bills Over Time




*Notes:* Figure shows the number of environmental bills voted on and the share that passed a simple majority in the House of Representatives from 1990 to 2000. Bill count is shown in blue bars (left axis) and share passed is shown in the red line (right axis).

Next, Table A.1 shows correlations between the tariff changes resulting from NAFTA and various district characteristics, as these motivate the inclusion of characteristic-trends in our analysis. District characteristics include various 1990 demographics, taken from Adler (2021), covering the share of the district's population aged 65 and above, the share black, the share born outside the US, and the share living in a rural area. We also include district median incomes and share of employment in farming, both from Adler, and the share of employment in manufacturing, taken from the adjusted county business patterns (CBP) database developed by Eckert et al. (2020). The first column shows correlations for the US tariff change and the second column shows correlations for the Mexican tariff change, with p-values in brackets.

Table A.1: Correlations Between District Tariff Changes and 1990 Characteristics

| | (1) US Tariff Change | (2) Mexican Tariff Change |
|---|---|---|
| Pop. Share - Above 65 | 0.036 | -0.089 |
| | [0.453] | [0.065] |
| Pop. Share - Black | 0.138 | -0.104 |
| | [0.004] | [0.031] |
| Pop. Share - Foreign | -0.064 | -0.011 |
| | [0.186] | [0.813] |
| Pop. Share - Rural | 0.119 | -0.026 |
| | [0.013] | [0.583] |
| Median Income | -0.408 | 0.111 |
| | [0.000] | [0.021] |
| Emp. Share - Farm | 0.101 | -0.142 |
| | [0.037] | [0.003] |
| Emp. Share - Manuf. | 0.555 | 0.347 |
| | [0.000] | [0.000] |

*Notes*: Table shows pairwise correlations between the districts tariff change as a result of NAFTA and various district characteristics. All district characteristics are measured for 1990. Column (1) shows correlations for the US import tariff change, while Column (2) shows correlations for the Mexican tariff change. Rows one through four show correlations between tariff changes and the share of the population aged 65 or older, black, born outside the US, and living in a rural area, respectively. Row five shows correlations between tariff changes and district median incomes. Rows six and seven show correlations between tariff changes and the share of the workforce employed in farming and manufacturing, respectively. The p-value on each correlation is shown in brackets.

The statistics in column (1) of Table A.1 indicate that the size of a district's US tariff reduction is positively correlated with the district's rural population share, share of the population that identifies as black, employment share in farming, and employment share in manufacturing, but negatively correlated with the district's income level. The second column indicates that the size of a district's Mexican tariff reduction is positively correlated with the district's employment share in manufacturing and income level, but negatively correlated with the share of the district's population above the age of 65, the share of the population that identifies as black, and the employment share in farming.

In addition, Table A.2 assess whether the magnitude of the NAFTA tariff changes vary systematically depending on pre-NAFTA political conditions. We do so by computing the correlation between each of the district's tariff changes (both Mexican and US) and the share of votes received by the Republican party in 1993. In the table, the first column reports the correlation for the US tariff changes, while the second column shows the correlation for the Mexican tariff changes, with p-values in brackets. As the table shows, the correlation between each tariff and the Republican vote share in the pre-NAFTA election are both small and statistically insignificant, suggesting changes in tariffs were unrelated to political conditions.

Lastly, Table A.3 shows correlations between each of the six policy preference variables used in our analysis of policy preferences (Table 7). Of note is that views on environmental protection are positively correlated with all additional policy measures.

Table A.2: Correlations Between District Tariff Changes and 1992 Republican Vote Share

| | (1) US Tariff Change | (2) Mexican Tariff Change |
|---|---|---|
| Republican Vote Share | -0.042 | 0.024 |
| | [0.383] | [0.616] |

*Notes*: Table shows the correlation between the district tariff change as a result of NAFTA and the vote share of the Republican party in the election prior to NAFTA. Column (1) shows the correlation for the US tariff changes. Column (2) shows the correlation for the Mexican tariff changes.

Table A.3: Correlations in Constituent Policy Views

| | (1) Env. Spending | (2) Welfare Spending | (3) Soc. Sec. Spending | (4) Crime Spending | (5) Pro-Choice | (6) Immig'n |
|---|---|---|---|---|---|---|
| Env. Spending | 1.000 | | | | | |
| Welfare Spending | 0.167 | 1.000 | | | | |
| Soc. Sec. Spending | 0.109 | 0.183 | 1.000 | | | |
| Crime Spending | 0.125 | 0.061 | 0.158 | 1.000 | | |
| Pro-Choice | 0.091 | 0.016 | -0.028 | -0.038 | 1.000 | |
| Immig'n | 0.049 | 0.084 | 0.001 | -0.032 | 0.026 | 1.000 |

*Notes*: Table shows correlations between each of the respondent policy views assessed in Table 5. Voter policy views are taken from the American National Election Studies survey.

That is, respondents who believe the federal government should spend more on environmental protection are more likely to believe the government should spend more on welfare, social security, and crime, and are more likely to be pro-choice and favor increased immigration.

# Appendix B Robustness Tests

We probe the robustness of our main results along six main dimensions.

First we examine whether our estimates of the effects of reductions in US import tariffs are solely capturing the effects of increased import competition from Mexico, or if they are also capturing the effects in reductions in the cost of importing intermediate inputs from Mexico. As we noted in the main text, there is reason to believe this may be the case as there was significant trade in intermediate goods between the US and Mexico prior to NAFTA. To examine this possibility, we re-estimate the specifications presented in Table 2 but replace reductions in district level US import tariffs with the change in each district's Effective Rate of Protection (ERP) (Corden, 1966):

$$\Delta ERP_r = \sum_i \left[\frac{l_{ir,90}}{l_{r,90}}\right] [ERP_{i,99} - ERP_{i,93}] \tag{B.1}$$

where $ERP_{i,t} = [Tariff_{i,t}^{USA} - \sum_j \alpha_{ij} Tariff_{i,t}^{USA}]/[1 - \sum_j \alpha_{ij}]$ and $\alpha_{ij}$ is industry $j$'s input share in the production of output from industry $i$. Because $\Delta ERP_r$ captures the net effect of lowering tariffs on both output and intermediate inputs, this exercise allows us to determine if our main estimates are capturing the effects of changes in tariffs on intermediate inputs.

The results from this exercise are presented in Table B.4. As these estimates show, accounting for the effects of reductions in tariffs in intermediate inputs has a modest effect on the magnitude of our empirical estimates; the estimates reported in Table B.4 are smaller in absolute value than the corresponding estimates reported in Table 2 of the main text. However, the estimates reported in Table B.4 are not statistically distinguishable from those reported in Table 2 of the main text at conventional levels, suggesting that our main estimates are primarily capturing the effects of changes in import competition created by reductions in tariffs on output.

Table B.4: The Effects of NAFTA on House Roll Call Votes - Effective Rate of Protection

| | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| $\Delta ERP_r$ x $\text{Post}_t$ | -0.088 | | -0.087 | -0.087 | -0.103 | -0.098 |
| | (0.026) | | (0.027) | (0.027) | (0.029) | (0.030) |
| $\Delta \tau_r^{Mex}$ x $\text{Post}_t$ | | -0.010 | -0.008 | -0.008 | -0.013 | -0.016 |
| | | (0.011) | (0.010) | (0.012) | (0.008) | (0.009) |
| CAA Trends | | | | X | | X |
| Charac. Trends | | | | | X | X |
| $R^2$ | 0.36 | 0.36 | 0.36 | 0.36 | 0.38 | 0.38 |
| Obs. | 50322 | 50322 | 50322 | 50322 | 50322 | 50322 |

*Notes*: Table reports estimates of the effects of reductions the Effective Rate of Protection in the US and Mexican tariffs on roll call votes on environmental bills in the House of Representatives between 1990 and 2000. The dependent variable in all regressions is an indicator for whether the roll call vote cast by a representative on a particular bill is pro-environment. All regressions include congressional district and year fixed effects. Column (1) shows the results of a simple difference-in-difference regression of the reduction in US import tariffs only. Column (3) shows the results of a simple difference-in-difference regression of the reduction in Mexican tariffs only. Column (3) shows the results of a simple difference-in-difference regression jointly estimating the effects of both tariff changes. Column (4) adds controls for the effects of the Clean Air Act with initial district non-attainment status by year fixed effects. Column (5) adds district baseline characteristic by year fixed effects. Column (6) corresponds to our baseline specification, and includes all additional controls and fixed effects. Standard errors two-way clustered by state and bill are shown in parentheses.

In our second robustness exercise, we examine other potential explanations for our results. The results of the first such exercise are reported in the nine columns of Table B.5. In columns (1) and (2) we examine whether our estimates are capturing the effects of ongoing changes in tariffs as the result of other relevant trade agreements, with column (1) addressing ongoing tariff changes due to the Canada-US Free Trade Agreement (CUSFTA) and column (2) addressing multilateral trade negotiations as part of the General Agreement on Tariffs and Trade. In column (1), to flexibly control for the effects of CUSFTA, we incorporate into our baseline regression an interaction between the change in both the district's Canadian and US tariffs resulting from CUSFTA that occurred after NAFTA's implementation (that is, between 1994 and 2000) with a post-NAFTA indicator.[1] In column (2), we include an interaction between the change in the

[1]We follow the same procedure outlined in Section 2 of the main text, and create the district's exposure

Most Favored Nation (MFN) tariffs that occurred after NAFTA's implementation with a post-NAFTA indicator. In column (3) we account for each district's exposure to trade with China, to ensure our results are capturing the effects of the China shock (Autor et al., 2013). We do so by controlling for the natural logarithm of the district's imports from and exports to China in each year.[2] In column (4) we supplement our baseline specification with indicators that reflect whether a given representative is a member of the majority party in the House of Representatives and Senate, or whether the representative's party affiliation aligns with the party of the President. We include these controls to account for differential voting incentives that may arise depending on who controls the Senate and Presidency. In column (5), we include bill fixed effects to ensure that we are not capturing idiosyncratic aspects of specific bills. In column (6), we include district by election-year fixed effects to account for the possibility of differential roll call voting behavior in election years. In column (7), we include Census Division by year fixed effects to ensure our results are not capturing differential trends across broadly defined regions. In column (8), we control for each district's share of workers employed in blue-collar jobs by including baseline blue-collar worker share by year fixed effects to ensure our results are not capturing the effects of industries already on the decline prior to NAFTA (see, e.g. Hakobyan and McLaren (2016)). Column (9) controls for all of these additional factors simultaneously.

As the estimates reported in Table B.5 show, our main findings are highly robust. The estimates reported in the table are similar in magnitude to those from our baseline specification, indicating that our baseline estimates are not capturing the effects of other factors.[3]

In our third robustness exercise, we examine the possibility that our results are capturing the effects of time-varying changes in political conditions. We omitted such changes from our baseline specifications as NAFTA exposure is uncorrelated with pre-NAFTA political conditions, as shown in Online Appendix A. For completeness sake, however, we report the results from controlling for such differential trends in Table B.6. We adopt three approaches to account for the possibility of differential trends based on initial political conditions. First, in column (1) we include an interaction between the Democratic party's initial vote share in the district and year fixed effects. Second, in column (2), we include an interaction between the party that holds the district's seat in the first year the district enters our sample (1990 for most districts, and 1993 for the districts created following redistricting) and a year fixed effect. Third, in column (3), we include both additional controls. As the estimates reported in Table B.6 show, accounting for a district's party of representation or voting patterns prior to NAFTA causes no

to CUSFTA as a weighted average of each industry's tariff changed, using district employment shares as weights.

[2]Similar to the approach taken by Autor et al., we construct measures of district imports and exports by allocating industry trade flows to the district level using the initial share of district industry employment in national industry employment as weights.

[3]Controlling for CUSFTA tariff changes appears to substantially increase the estimated effects of the US tariff changes (Column (1)), while controlling for MFN tariff changes appears to substantially increase the estimated effects of the Mexican tariff changes (Column (2)), both of which suggest our baseline regression may be underestimating NAFTA's effects on RCVs. However, these estimates are not statistically different from our baseline estimates at conventional levels.

Table B.5: The Effects of NAFTA on House Roll Call Votes - Robustness Tests

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
|---|---|---|---|---|---|---|---|---|---|
| $\Delta\tau_r^{USA}$ x $\text{Post}_t$ | -0.258 | -0.119 | -0.149 | -0.122 | -0.152 | -0.148 | -0.093 | -0.141 | -0.183 |
| | (0.086) | (0.057) | (0.047) | (0.043) | (0.048) | (0.046) | (0.055) | (0.046) | (0.086) |
| $\Delta\tau_r^{Mex}$ x $\text{Post}_t$ | -0.009 | -0.058 | -0.016 | -0.017 | -0.015 | -0.016 | -0.023 | -0.016 | -0.094 |
| | (0.017) | (0.039) | (0.009) | (0.010) | (0.009) | (0.009) | (0.010) | (0.009) | (0.042) |
| CUSFTA | X | | | | | | | | X |
| MFN | | X | | | | | | | X |
| China Shock | | | X | | | | | | X |
| Alignm. Vars. | | | | X | | | | | X |
| Bill FEs | | | | | X | | | | X |
| Dist.-Elec. FEs | | | | | | X | | | X |
| Cen. Div-Year FEs | | | | | | | X | | X |
| Blue Collar | | | | | | | | X | X |
| $R^2$ | 0.38 | 0.38 | 0.38 | 0.39 | 0.44 | 0.38 | 0.38 | 0.38 | 0.46 |
| Obs. | 50322 | 50322 | 50322 | 50322 | 50322 | 50322 | 50322 | 50322 | 50322 |

*Notes*: Table shows results of the NAFTA tariff cuts on roll call votes in the House of Representatives. The dependent variable in all regressions is an indicator for whether the roll call vote cast by a representative on a particular bill is pro-environment. All regressions include district and year fixed effects, and district baseline characteristic and CAA non-attainment status by year fixed effects. Column (1) includes interactions between the change in both the district's Canadian and US tariffs resulting from the Canada-US Free Trade Agreement (CUSFTA) that occurred after NAFTA's implementation (that is, between 1994 and 2000) with a post-NAFTA indicator. In column (2), we include an interaction between the change in the World Trade Organization's Most Favored Nation (MFN) tariffs that occurred after NAFTA's implementation with a post-NAFTA indicator. Column (3) controls for the natural log of each district's exports to and imports from China. Column (4) adds variables capturing the representatives alignment with the party in power in the House, Senate, and Presidency. Column (5) includes bill fixed effects. Column (6) adds district by election-year fixed effects. Column (7) includes census division by year fixed effects. Column (8) adds the district's share of workers in blue collar sectors to the set of baseline characteristic by year fixed effects. Column (9) adds to the baseline regression all additional controls and fixed effects from Columns (1) to (8).

meaningful change in our estimates of NAFTA's effects on RCVs.

Table B.6: The Effects of NAFTA on House Roll Call Votes with Initial Political Conditions

| | (1) | (2) | (3) |
|---|---|---|---|
| $\Delta\tau_r^{USA}$ x $\text{Post}_t$ | -0.146 | -0.146 | -0.146 |
| | (0.045) | (0.046) | (0.045) |
| $\Delta\tau_r^{Mex}$ x $\text{Post}_t$ | -0.015 | -0.015 | -0.015 |
| | (0.009) | (0.009) | (0.009) |
| Init. Party | X | | X |
| Democrat Share | | X | X |
| $R^2$ | 0.38 | 0.38 | 0.38 |
| Obs. | 50322 | 50322 | 50322 |

*Notes*: Table shows results of the reductions in US import tariffs and Mexican tariffs on roll call votes on environmental bills in the House of Representatives between 1990 and 2000, controlling for pre-NAFTA political conditions. Column (1) includes an interaction between the Democratic party's vote share in the first year the district enters our sample and a year fixed effect. Column (2) includes an interaction between the party that holds the district's seat in the first year the district enters our sample and a year fixed effect. Column (3) includes both controls. The dependent variable is an indicator for whether the roll call vote cast by a representative on a particular bill is pro-environment. The regression includes district Clean Air Act and baseline characteristic by year fixed effects, and district and year fixed effects. Standard errors two-way clustered by state and bill are shown in parentheses.

The fourth dimension along which we probe our main results is to ensure they do not reflect differential trends in outcomes across districts. We do so by estimating the following event-study version of our baseline specification:

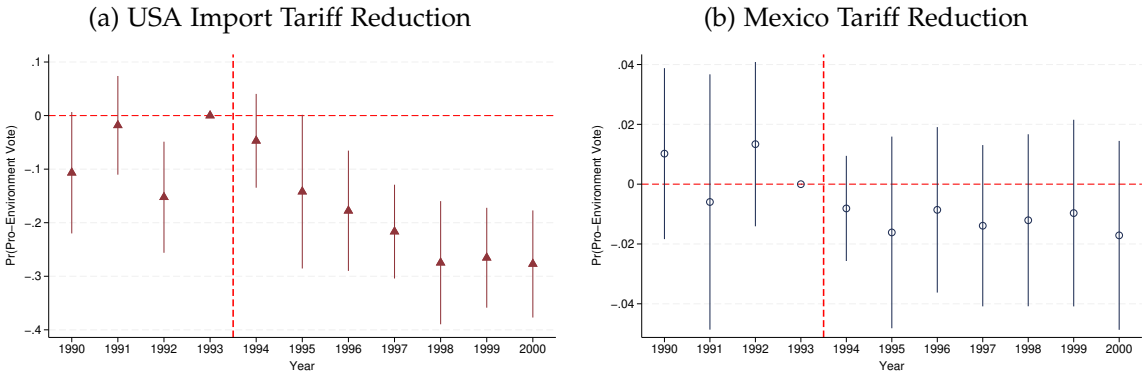
$$y_{vrt} = \beta_0 + \sum_{k=-m}^{M} \beta_{USA}^k \left[\tau_r^{USA} \times \mathbf{1}(t=k)\right] + \sum_{k=-m}^{M} \beta_{Mex}^k \left[\tau_r^{Mex} \times \mathbf{1}(t=k)\right] + \lambda_r + \psi_t + e_{vrt} \tag{B.2}$$

where the regression coefficients $\beta_{USA}^k$ and $\beta_{Mex}^k$ measure the effect of the changes in US import tariffs and Mexican tariffs, respectively, in the $m$ years before to the $M$ years after NAFTA, and all other variables are defined as before. If, as we have assumed, there are no other factors aside from NAFTA driving differential trends across districts, then we should observe $\hat{\beta}_{USA}^k = 0$ and $\hat{\beta}_{Mex}^k = 0$ for $m = \{1990, 1991, 1992\}$.

The results of this analysis are displayed in the two panels of Figure B.2.[4] Panel (a) depicts our estimates of $\beta_{USA}^k$, while panel (b) depicts our estimates of $\beta_{Mex}^k$. In both cases the associated 95% confidence intervals constructed using standard errors that are two-way clustered by state and bill are plotted around the estimates.

[4]The corresponding point estimates and standard errors are available from the authors on request.

Figure B.2: House Roll Call Vote Event Study




*Notes:* Figure shows coefficient estimates from a difference-in-difference event study estimating the effects of the NAFTA tariff reductions on roll call votes in the House of Representatives. Panel (a) shows estimates of the effects of US import tariff reductions and panel (b) shows estimates of the effects of Mexican tariff reductions. The dependent variable is an indicator for whether the roll call vote cast by a representative on a particular bill is pro-environment. The regression includes district, year, and district-by-election year fixed effects, as well as initial CAA non-attainment status and district baseline characteristic by year fixed effects. The year prior to NAFTA, 1993, is the omitted category. 95% confidence intervals from standard errors two-way clustered by state and bill are plotted around the coefficient estimates.

The coefficients plotted in the figure suggest that our baseline estimates are not simply capturing pre-existing differences in trends across districts, as $\hat{\beta}^k_{USA}$ and $\hat{\beta}^k_{Mex}$ are, for the most part, small and not statistically different from zero prior to 1994.[5] Moreover, the coefficient estimates displayed in Panel (a) indicate that the effect of the US import tariff reduction increased in magnitude between 1994 and 1998, suggesting that NAFTA's effect on environmental voting grew over time. This potentially reflects the fact that many of NAFTA's tariff reductions were phased-in over our period of study.[6]

For our fifth robustness exercise, we examine whether our results are robust to several alternative samples. First, we reproduce the analysis in Table 2 of the main text dropping any environmental bills related to fossil fuels, as these bills may be treated differently compared to legislation on other environmental issues. These results, shown in Table B.7, indicate that the effects of NAFTA on non-fossil fuel related environmental bills are very similar to our main estimates.

We also produce an event study using this sample of bills by estimating Equation (B.2). The results of the event study are shown in Figure B.3, with panel (a) showing the estimates for the US tariff change and panel (b) showing the estimates for the Mexi-

[5]One notable exception is that $\hat{\beta}^{1992}_{USA}$ is negative and statistically significant. This is caused by environmental bills that regulate fossil fuels, of which there were an unusually large number in 1992. This produces this effect for two reasons. First, fossil fuel-related bills in our dataset receive less support than other environmental bills (42% vs. 50% pro-environment). Second, there is a negative correlation between a district's tariff change and their support for fossil fuel-related bills prior to NAFTA. In Figure B.3, we show that dropping the 18 bills related to fossil fuels from our analysis eliminates any significant estimates prior to NAFTA, but leaves our main results unchanged.

[6]Over 50% of US tariffs on Mexican imports and 31% of Mexican tariffs on US imports were removed immediately upon NAFTA's implementation, while the majority of the remaining tariffs were removed according to predetermined schedules within ten years (Kowalczyk and Davis, 1998).

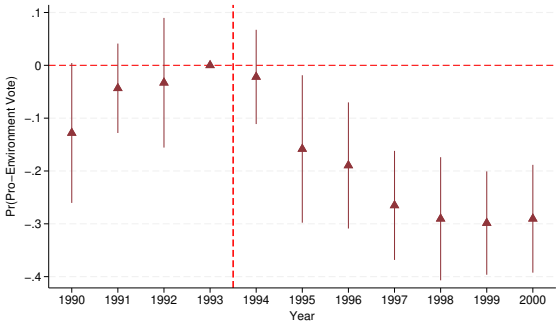
Table B.7: The Effects of NAFTA on Non-Fossil Fuel Roll Call Votes

| | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| $\Delta\tau_r^{USA}$ x $\text{Post}_t$ | -0.159 | -0.158 | -0.185 | -0.182 |
| | (0.043) | (0.043) | (0.049) | (0.049) |
| $\Delta\tau_r^{Mex}$ x $\text{Post}_t$ | -0.007 | -0.007 | -0.012 | -0.015 |
| | (0.011) | (0.013) | (0.009) | (0.009) |
| CAA Trends | | X | | X |
| Charac. Trends | | | X | X |
| $R^2$ | 0.41 | 0.41 | 0.43 | 0.43 |
| Obs. | 42466 | 42466 | 42466 | 42466 |

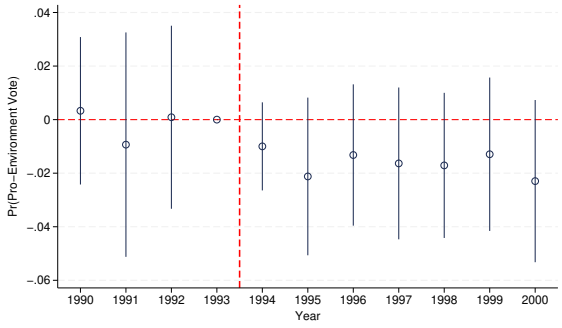
*Notes*: Table shows results of the reductions in US import tariffs and Mexican tariffs on roll call votes on environmental bills in the House of Representatives between 1990 and 2000, omitting any bills that pertain to fossil fuels. The dependent variable in all regressions is an indicator for whether the roll call vote cast by a representative on a particular bill is pro-environment. All regressions include congressional district and year fixed effects. Column (1) shows the results of a simple difference-in-difference regression. Column (2) controls for the effects of the Clean Air Act. Column (3) includes district baseline characteristics by year fixed effects. Column (4) is the baseline analysis, which includes all additional controls and fixed effects. Standard errors two-way clustered by state and bill are shown in parentheses.

Figure B.3: House Roll Call Vote Event Study Without Energy Bills

(a) USA Import Tariff Reduction




(b) Mexico Tariff Reduction




*Notes:* Figure shows coefficient estimates from a difference-in-difference event study estimating the effects of the NAFTA tariff reductions on roll call votes in the House of Representatives, omitting any bills that pertain to fossil fuels. Panel (a) shows estimates of the effects of US import tariff reductions and Panel (b) shows estimates of the effects of Mexican tariff reductions. The dependent variable is an indicator for whether the roll call vote cast by a representative on a particular bill is pro-environment. The regression includes district and year fixed effects, and district baseline Clean Air Act non-attainment status and characteristic by year fixed effects. The year prior to NAFTA, 1993, is the omitted category. 95% confidence intervals from standard errors two-way clustered by state and bill are plotted around the coefficient estimates.

can tariff change. Omitting fossil fuel related bills lends further confidence to our results. The US tariff reductions produce no significant change in RCVs prior to NAFTA, and cause a significant reduction in pro-environmental voting after NAFTA's introduction.

We then examine five additional samples. The results of these regressions are shown in Table B.8. In column (1) we restrict our sample to the years 1993 onward to ensure that our baseline estimates are not being driven by the district reapportionment that occurred following the 1990 census. In column (2) we restrict our sample to exclude

bills where the issue classification includes "other" to ensure that our estimates are not potentially capturing voting on other issues that have been included on environmental bills. In column (3) we omit the twenty-four congressional districts that experienced an increase in Mexican tariffs over our sample period, as the political conditions in these districts may be systematically different from the rest of the country. In column (4) we restrict our sample to omit bills that are subject to multiple roll call votes, as the votes for these bills may be subject to different incentives than other votes in our sample. In column (5) we expand our sample to treat RCV abstentions as negative votes following the classification scheme used by the LCV.

Table B.8: The Effects of NAFTA on House Roll Call Votes - Alternative Samples

| | (1) | (2) | (3) | (4) | (5) |
|---|---|---|---|---|---|
| $\Delta\tau_r^{USA}$ x $\text{Post}_t$ | -0.214 | -0.141 | -0.165 | -0.122 | -0.136 |
| | (0.047) | (0.044) | (0.044) | (0.047) | (0.047) |
| $\Delta\tau_r^{Mex}$ x $\text{Post}_t$ | -0.011 | -0.014 | -0.031 | -0.013 | -0.016 |
| | (0.011) | (0.008) | (0.033) | (0.009) | (0.009) |
| $R^2$ | 0.42 | 0.35 | 0.39 | 0.36 | 0.35 |
| Obs. | 41694 | 39828 | 47702 | 41114 | 52312 |

*Notes*: Table shows results of the NAFTA tariff cuts on roll call votes in the House of Representatives for various samples. The dependent variable in all regressions is an indicator for whether the roll call vote cast by a representative on a particular bill is pro-environment. All regressions include district and year fixed effects, and district baseline characteristic and CAA non-attainment status by year fixed effects. Column (1) restricts the sample to years after redistricting (1993-2000). Column (2) omits any bills that may address non-environmental issues (in addition to environmental issues). Column (3) omits any districts that experienced an increase in average export tariffs. Column (4) omits any bills that are subject to multiple roll call votes. Column (5) includes abstentions and classifies them as "negative" votes. Standard errors two-way clustered by state and bill are shown in parentheses.

As the estimates reported in Table B.8 show, all restricted samples produce estimates that are not statistically distinguishable from those in our baseline specification, which suggests that our preferred estimates are not capturing the effects of redistricting, particular characteristics of certain bills and districts, or the LCV's treatment of abstentions.

Lastly, as our final robustness exercise, we examine whether the effect of NAFTA's tariff changes varies depending on the type of legislation being considered, given that legislator incentives may change throughout the voting process. Here, we separately estimate NAFTA's effects on three different types of RCVs: those on the final passage of a bill, those on a proposed amendment to a bill, and those on a motion. These results, shown in Table B.9, indicate that the effects of NAFTA on RCVs is remarkably consistent across each of these three types of RCVs, suggesting our estimates are not masking underlying heterogeneity in voting responses.

# Appendix C The Introduction of New Environmental Bills

As discussed in the main text, a potential issue with our estimates is that they may be biased due to a selection effect created by a NAFTA-induced change in the set of bills that appear before Congress. That is, if NAFTA systematically changed the set of

Table B.9: The Effects of NAFTA on RCVs, by Vote Type

| | Final Bill (1) | Amendment (2) | Motion (3) |
|---|---|---|---|
| $\Delta\tau_r^{USA}$ x $\text{Post}_t$ | -0.166 | -0.155 | -0.163 |
| | (0.083) | (0.047) | (0.053) |
| $\Delta\tau_r^{Mex}$ x $\text{Post}_t$ | -0.004 | -0.018 | 0.001 |
| | (0.010) | (0.009) | (0.013) |
| CAA Trends | X | X | X |
| Charac. Trends | X | X | X |
| $R^2$ | 0.45 | 0.37 | 0.63 |
| Obs. | 12488 | 36166 | 1648 |

*Notes*: Table shows results of the reductions in US import tariffs and Mexican tariffs on roll call votes on environmental bills in the House of Representatives between 1990 and 2000. Column (1) restricts the sample to RCVs on bill passage. Column (2) restricts the sample to RCVs on proposed bill amendments. Column (3) restricts the sample to RCVs on motions. The dependent variable in all regressions is an indicator for whether the roll call vote cast by a representative on a particular bill is pro-environment. All regressions include congressional district and year fixed effects, controls for the effects of the Clean Air Act, and district baseline characteristics by year fixed effects. Standard errors two-way clustered by state and bill are shown in parentheses.

environmental bills introduced in the House, then comparing roll call votes before and after NAFTA would misrepresent NAFTA's effect on RCVs. Though we have strong reason to believe this concern is minor in our setting, as we discuss in Section 3.1 of the main paper, here we examine this issue directly by estimating NAFTA's effects on the likelihood that a congressperson introduced a new environmental bill, as well as the complexity of the environmental bills introduced (measured by number of committee referrals). If bill selection is an important concern, then we should find a change in bill proposals or complexity by legislators more exposed to NAFTA's tariff reductions.

To perform this exercise, we collect data on all bill proposals in the House between 1990 and 2000 from the Congressional Bills Project data of Adler and Wilkerson (2021). The Congressional Bills Project records information on all bill proposals to the House between 1947 and 2008. The dataset includes information on the bill's sponsor, committee referrals, and a categorization of it's main topic.[7] We use this information to construct a district-level panel capturing the introduction of new bills by the district's representative. We use this data to collect all bills that are related to the environment, and then construct two measures for each district-year: an indicator of whether the district's representative introduced at least one environment-related bill that year and an indicator of whether any of their environmental bills were referred to multiple committees.

Before discussing our analysis, we first describe our approach to measuring bill complexity. While a full examination of bill content is beyond the scope of this paper, we examine a simple measure of bill complexity: the number of committees to which a bill has been referred. After a bill is introduced in the House, it must be referred to com-

[7]The dataset categorizes bills into 23 different topic areas, using the topic definitions from the Comparative Agendas Project. The topic list is available at: `http://www.comparativeagendas.net/pages/master-codebook`.

mittee for further assessment, before potentially returning to the House floor for a roll call vote. Bills may be referred to one or more committees for assessment. The ability to refer bills to multiple committees is a relatively recent change to congressional rules; it was introduced to the House in 1975 to both aid in assessing complex policy issues and to encourage inter-committee cooperation on jurisdictional conflicts (Davidson et al., 1988). Thus, bills assigned to multiple committees should, on average, be more complex than single-committee bills. We use this logic to examine whether NAFTA affected the complexity of new environmental bills.

With this data, we then estimate a generalized difference-in-difference regression analogous to that used in our main analysis by estimating the following regression:

$$b_{rt} = a_0 + a_{USA}\left[\Delta\tau_r^{USA} \text{ x } \text{Post}_t\right] + a_{Mex}\left[\Delta\tau_r^{Mex} \text{ x } \text{Post}_t\right] + \lambda_r + \psi_t + e_{rt}, \tag{C.3}$$

where $r$ and $t$ index house districts and years, respectively, and $b_{rt}$ is either the new bill indicator or multiple referral indicator. In Equation (C.3), all other variables are as in Equation (1), and $a_{USA}$ and $a_{Mex}$ are our estimates of the effects of a 1 pp reduction in US and Mexican tariffs, respectively. Lastly, we cluster standard errors by state.

The results of this analysis are presented in the two panels of Table C.1. In Panel (a), the dependent variable is the indicator of whether the district's representative introduced at least one environmental bill in a particular year. The sample for this analysis includes all district-years. In Panel (b), the dependent variable is an indicator for whether the district's representative introduced an environmental bill that was referred to multiple committees that year. The sample for this analysis only includes district-years that introduced at least one environmental bill. Each panel shows results of four specifications, each of which includes a different set of controls, as indicated by the table.

The results in Panel (a) of Table C.1 show reductions in both US import tariffs and Mexican tariffs did not significantly impact the introduction of environmental bills. For example, our baseline estimates (column (4)), indicate that a 1 pp reduction in US import tariffs reduced this likelihood by 8.4 pp. Not only is this estimate not statistically different from zero, but it is economically small as well. Given the average reduction in US import tariffs across districts is 0.25 pp, this suggests that NAFTA reduced the likelihood of introducing a new environmental bill by 2.1 pp.[8]

The results in Panel (b) indicate that neither the US nor Mexican tariff reductions had a measurable effect on committee referrals. For example, our baseline estimate (column (8)) shows that reductions in US import tariffs caused a small, but statistically insignificant, decrease in the likelihood that a district's representative had an environmental bill referred to multiple committees. On average, US import tariff reductions reduced the likelihood of a multiple bill referral by less than 1 pp among district-years with at least one environmental bill.[9] As approximately 40% of district-years that introduce an environmental bill have at least one referred to multiple committees in our sample, the effects of both the US and Mexican tariff reductions on multiple referrals appear to be relatively

[8]Note that representatives in 22% of district-years introduced a new environmental bill in our sample.

[9]This statistic is computed by multiplying the point estimate in Column (4) of Table C.1 by the average reduction in district import tariffs.

Table C.1: NAFTA and the Introduction and Complexity of Environmental Bills

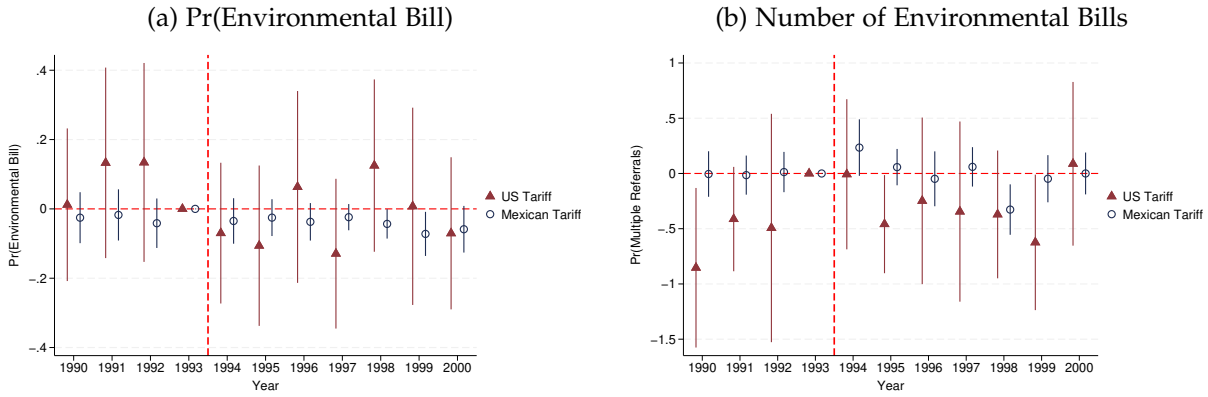
| | Panel (a): Pr(Environmental Bill) | | | |
|---|---|---|---|---|
| | (1) | (2) | (3) | (4) |
| $\Delta\tau_r^{USA}$ x $\text{Post}_t$ | -0.035 | -0.036 | -0.085 | -0.084 |
| | (0.048) | (0.048) | (0.066) | (0.066) |
| $\Delta\tau_r^{Mex}$ x $\text{Post}_t$ | -0.012 | -0.015 | -0.021 | -0.023 |
| | (0.021) | (0.022) | (0.028) | (0.028) |
| CAA Trends | | X | | X |
| Charac. Trends | | | X | X |
| $R^2$ | 0.24 | 0.25 | 0.27 | 0.27 |
| Obs. | 4767 | 4767 | 4767 | 4767 |
| | Panel (b): Pr(Multiple Referral) | | | |
| | (5) | (6) | (7) | (8) |
| $\Delta\tau_r^{USA}$ x $\text{Post}_t$ | -0.135 | -0.146 | -0.014 | -0.032 |
| | (0.138) | (0.135) | (0.174) | (0.197) |
| $\Delta\tau_r^{Mex}$ x $\text{Post}_t$ | 0.056 | 0.067 | 0.005 | 0.015 |
| | (0.035) | (0.034) | (0.030) | (0.031) |
| CAA Trends | | X | | X |
| Charac. Trends | | | X | X |
| $R^2$ | 0.37 | 0.38 | 0.45 | 0.46 |
| Obs. | 952 | 952 | 952 | 952 |

*Notes*: Table shows results of the NAFTA tariff cuts on the introduction of new bills pertaining to the environment, energy, or public lands in the House of Representatives. The dependent variable in Panel (a) is an indicator of whether the district's representative introduced a new bill in a particular year. The dependent variable in Panel (b) is an indicator of whether an environmental bill introduced by the district's representative in a given year was referred to multiple committees, estimated on the sample of district-years for which the district's representative sponsored a new environmental bill. All regressions include congressional district and year fixed effects. Column (1) shows the results of a simple difference-in-difference regression. Column (2) controls for the effects of the Clean Air Act. Column (3) includes district baseline characteristic by year fixed effects. Column (4) is the baseline analysis, which includes all additional controls and fixed effects. Standard errors clustered by state are shown in parentheses.

small. This suggests that NAFTA did little to alter the complexity of the environmental bills introduced in the House, as measured by committee referrals.

To assess the robustness of the results presented in Table C.1, we estimate an event study variant of Equation (C.3) for both dependent variables, adopting our baseline specification (columns (4) and (8) in Table C.1). Coefficient estimates and associated 95% confidence intervals from both event studies are shown in Figure C.1. In Panel (a), the dependent variable is the indicator of whether the district's representative introduced at least one environmental bill in a particular year. In Panel (b), the dependent variable is our multiple-committee referral indicator. Both event study estimates show no meaningful pattern for either the US or Mexican tariff reductions, further suggesting that bill selection is not of material importance in our setting.

Figure C.1: Bill Selection Event Study




*Notes:* Figure shows coefficient estimates from an study estimating the effects of the NAFTA tariff cuts on the introduction of new bills pertaining to the environment, energy, or public lands in the House of Representatives. The dependent variable in Panel (a) is an indicator of whether the district's representative introduced a new bill in a particular year. The dependent variable in Panel (b) is an indicator of whether an environmental bill introduced by the district's representative in a given year was referred to multiple committees, estimated on the sample of district-years with at least one environmental bill. In each panel, coefficient estimates and 95% confidence intervals are shown for US tariffs (in blue) and Mexican tariffs (in red). All regressions include congressional district and year fixed effects, and district characteristic and CAA non-attainment by year fixed effects. The year prior to NAFTA, 1993, is the omitted category. 95% confidence intervals from standard errors clustered by state are plotted around the coefficient estimates.

# Appendix D An Alternative Estimator

As noted in the main text, one additional concern with our baseline estimates is that they may be biased due to the presence of systematic differences in treatment effects across groups or time. This potential concern arises because we have implemented our research design using a two-way fixed effect estimator. However, as shown by de Chaisemartin and D'Haultfouille (2022), if there are differences in treatment effects across groups or time, then the treatment effect estimates returned from such estimators are a weighted average of these underlying heterogeneous effects, where the weights may be negative. Thus, one may be concerned that our finding of a negative effect of the US import tariffs on RCVs is simply a spurious result due to the presence of negative weights in our two-way fixed effect regression. To address this concern, we implement our research design using the $DID_l$ estimator proposed by de Chaisemartin and D'Haultfouille (2022), which is robust to the presence of treatment-heterogeneity and dynamic treatment effects.

The results of this exercise are reported in Table D.1, which displays estimates from our main empirical specification (Equation (1)) as implemented by the $DID_l$ estimator. We report coefficient estimates from four specifications. As in Table 2 of the main paper, column (1) reports estimates which only includes district and year fixed effects. Column (2) adds initial district CAA non-attainment status by year fixed effects. Column (3) includes initial district-characteristic by year fixed effects. Finally, column (4), corresponds to our baseline specification which simultaneously includes initial district CAA non-attainment status and district-characteristic by year fixed effects. Given the nature of the $DID_l$ estimator, each specification reports estimates of the US import tariff's effects on the likelihood of casting a pro-environment RCV by year, controlling for Mexican import

Table D.1: The Effects of NAFTA on RCVs: An Alternative Estimator

| | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| $\Delta\tau_r^{USA}$ | | | | |
| x 1994 | -0.036 | -0.053 | -0.101 | -0.101 |
| | (0.015) | (0.018) | (0.035) | (0.035) |
| x 1995 | -0.093 | -0.123 | -0.218 | -0.218 |
| | (0.036) | (0.039) | (0.074) | (0.074) |
| x 1996 | -0.084 | -0.132 | -0.274 | -0.274 |
| | (0.030) | (0.044) | (0.104) | (0.104) |
| x 1997 | -0.091 | -0.161 | -0.364 | -0.364 |
| | (0.032) | (0.058) | (0.133) | (0.133) |
| x 1998 | -0.119 | -0.195 | -0.439 | -0.439 |
| | (0.033) | (0.068) | (0.163) | (0.163) |
| x 1999 | -0.109 | -0.200 | -0.483 | -0.483 |
| | (0.036) | (0.074) | (0.188) | (0.188) |
| x 2000 | -0.128 | -0.247 | -0.578 | -0.578 |
| | (0.033) | (0.086) | (0.217) | (0.217) |
| N | 35,929 | 35,929 | 35,929 | 35,929 |

*Notes*: Table shows results of NAFTA's US import tariff reduction on the likelihood of a pro-environment RCV, using de Chaisemartin and D'Haultfouille (2020)'s $DID_l$ estimator that is robust to treatment heterogeneity and dynamic treatment effects. Estimates for each year from 1994 to 2000 are shown. Results from three specifications are shown. Each regression includes district and year fixed effects. Column (1) has no controls, Column (2) adds initial Clean Air Act (CAA) non-attainment status by year fixed effects , Column (3) adds baseline characteristic by year fixed effects, and Column (4) includes initial CAA non-attainment status and baseline characteristic by year fixed effects. Standard errors are cluster-bootstrapped by state, using 300 repetitions. The table also shows the number of observations used in estimation (N).

tariff changes. In all cases, bootstrapped standard errors, clustered by state, are reported in parentheses.[10]

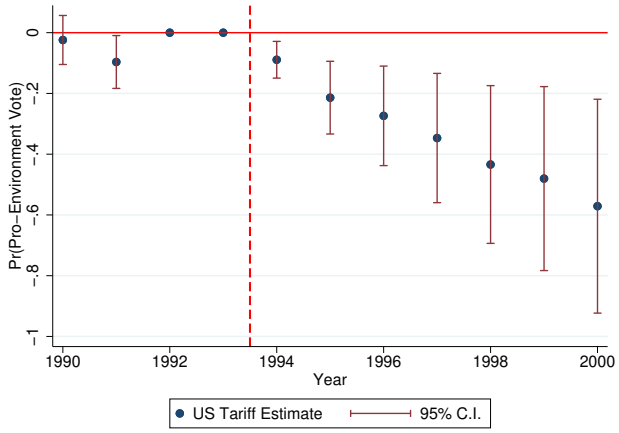
The estimates reported in Table D.1 are consistent with our main results. For our baseline specification (column (4)), a reduction in US import tariffs caused a statistically significant reduction in pro-environment RCVs in each year from 1994 to 2000. As in the event-study estimates reported in Figure B.2, the magnitude of the US import tariff change effect also increases over time. In addition, the estimates from the alternative specifications (columns (1)-(3)) all show a similar pattern, although the estimated magnitudes are smaller without the inclusion of initial district-characteristic by year fixed effects.

For our baseline specification, (column (4) in Table D.1), we also use the $DID_l$ estimator to perform an alternative event-study style placebo test for the presence of pre-existing differences in trends across the treated and control groups. This exercise uses $DID_l$ estimation to estimate treatment effects two or more years prior to treatment, omitting the year immediately prior to treatment. The results of this exercise are shown in Figure D.1, which shows the placebo estimates from 1990 to 1992 and the main treatment

[10]These standard errors are bootstrapped 300 times. We cluster by state rather than by state and bill as the $DID_l$ estimator does not allow for two-way clustering.

effect estimates from 1994 to 2000. For each estimate, a 95% confidence interval is displayed, produced from standard errors bootstrap-clustered by state with 300 repetitions.

Figure D.1: $DID_l$ Placebo Estimates




*Notes:* Figure shows results of NAFTA's US import tariff reduction on the likelihood of a pro-environment RCV, using de Chaisemartin and D'Haultfouille (2022)'s $DID_l$ estimator that is robust to treatment heterogeneity and dynamic treatment effects. Placebo treatment effect estimates from 1990 to 1992 and treatment effect estimates from 1994 to 2000 are shown with a 95% confidence interval. Standard errors are cluster-bootstrapped by state, using 300 repetitions.

The placebo estimates in Figure D.1 indicate that are main results are not simply due to pre-existing differential trends in RCVs, as they show no meaningful pattern prior to NAFTA. The placebo estimates are all relatively small in magnitude and are statistically indistinguishable from zero in 1990 and marginally significant in 1991.[11] This corroborates the results of the event study analysis presented in Figure B.2, which also indicated that pre-existing differences by trade-exposure are not an issue in our setting. As these placebo estimates are robust to the presence of dynamic treatment effects, they provide further support for our research design.

# Appendix E Additional Results

This section presents additional empirical results referenced in the main text. Section E.1 presents additional event study results, while Section E.2 presents other results.

## E.1 Event Study Results

This subsection presents additional event study results to complement the analysis presented in the main paper. To save space, coefficient estimates and 95% confidence intervals are shown for each event study.[12]

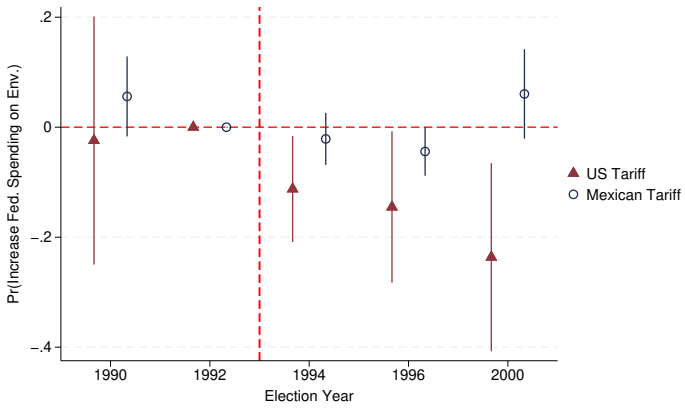
First, to complement the analysis presented in Section 3.2 that examines NAFTA's effects on the demand for environmental policy, we produce event studies for voters' stated views on environmental policy, and county economic and environmental conditions. The

[11] The estimate and standard error in 1992 is very small, which is why it appears omitted in the figure.

[12] Result tables are available upon request.

event study on voters' views on environmental policy is shown in Figure E.1. The dependent variable in this event study is an indicator for whether the respondent feels the federal government should increase spending on environmental protection. The regression includes initial district CAA non-attainment status and baseline-characteristic by year fixed effects, respondent and interviewer demographic-by-year and by-state fixed effects, and district and year fixed effects, with standard errors clustered by state. The omitted year in all regressions for both the US and Mexican tariffs is 1993, the year prior to NAFTA.[13] Estimates for US tariff changes are shown in blue; estimates for Mexican tariff changes are shown in red. As the figure shows, US and Mexican tariff changes prior to NAFTA had no significant effect on environmental policy views of respondents. Following NAFTA, US import tariff reductions reduced support for the environment, with the peak occurring in 2000. In contrast, Mexican tariff changes had no significant effect on environmental policy views post-NAFTA.

Figure E.1: Voter Environmental Policy View Event Study



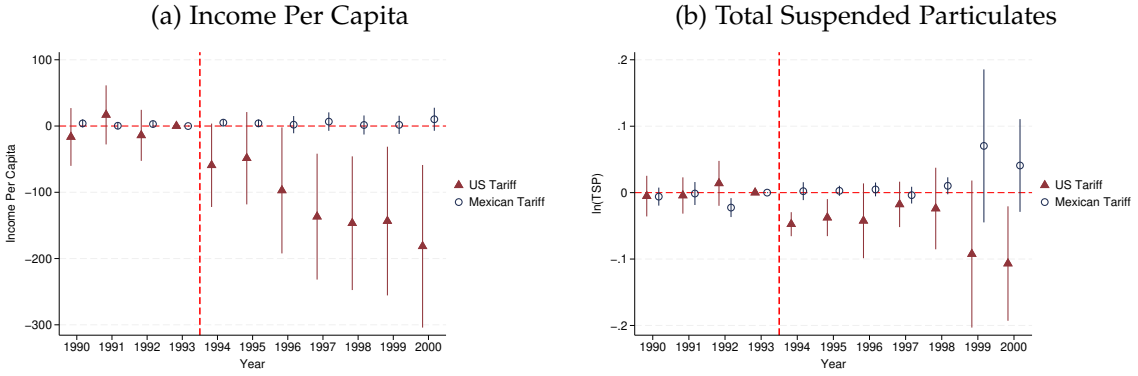

*Notes:* Figure shows coefficient estimates from a difference-in-difference event study estimating the effects of the NAFTA tariff reductions on views expressed on environmental policy stringency by voters between 1990 and 2000. Voter views are taken from the American National Election Studies survey. The dependent variable is an indicator of whether the survey respondent believes the federal government should increase spending on environmental protection. Coefficient estimates and 95% confidence intervals are shown for the US tariff reduction (in blue) and the Mexican tariff reduction (in red). The regression includes district and year fixed effects, district baseline characteristic and CAA non-attainment by year fixed effects, and respondent and interviewer demographics interacted with year and state fixed effects. Standard errors clustered by state are shown in parentheses. The year prior to NAFTA, 1993, is the omitted category. Standard errors are clustered by state. The year 1998 is omitted, as this question was not asked in that survey.

Second, Figure E.2 shows the results of event studies examining the effects of NAFTA on county income per capita (Panel (a)) and ambient total suspended particulate concentrations (Panel (b)) between 1990 and 2000. The dependent variable in Panel (a) is the county's per capita income, while the dependent variable in Panel (b) is the natural log of the county's average daily TSP concentration recorded over the year. Each regression includes initial county CAA non-attainment status and baseline-characteristic by year fixed effects, and county and year fixed effects, with standard errors clustered by state. The omitted year for both the US and Mexican tariffs is 1993, the year prior to NAFTA. Estimates for US tariff changes are shown in blue; estimates for Mexican tariff changes

[13]The relevant ANES question was not asked in 1998. Hence, that year is omitted from the regression.

are shown in red.

Figure E.2: County Economic and Environmental Conditions Event Studies




*Notes:* Figure shows coefficient estimates from a difference-in-difference event study estimating the effects of the NAFTA tariff reductions on county economic conditions and environmental quality. The dependent variable in Panel (a) is the county's average income per capita, while the dependent variable in Panel (b) is the natural log of the county's median daily ambient total suspended particulate concentration. In each panel, coefficient estimates and 95% confidence intervals are shown for the import shock (in blue) and the export shock (in red). Each regression includes county baseline characteristic and CAA non-attainment status by year fixed effects, and county and year fixed effects. The year prior to NAFTA, 1993, is the omitted category. Standard errors are clustered by state.

The results in Panel (a) of Figure E.2 show no significant effect of NAFTA on county incomes prior to 1994, and a stark reduction tied to import tariffs beginning in 1994 and persisting throughout the decade. The results in Panel (b) show no significant effect of the import tariff reductions prior to NAFTA's implementation, with a significant reduction in TSP following NAFTA, although the effect is not statistically significant in the years between 1996 and 1998. Panel (b) also shows no meaningful pattern with respect to the Mexican tariff reductions.

## E.2 Further Evidence of NAFTA's Effects on The Demand For Environmental Policy and Partisan Representation

This section presents additional results referenced in Section 4.4.1 of the main text.

### E.2.1 NAFTA and the Demand for Environmental Policy: Redux

As we note in the main text, the estimates presented in Table 6 indicate that changes in the voting behavior of incumbent Republicans explain close to half of NAFTA's effects on the formation of environmental policy in the US. The results presented in Table 4 suggest that such changes are due to these legislators responding to the demands of their constituents. However, those estimates capture the average effect of tariff changes across all affected districts, meaning that they need not necessarily reflect changes in the demands of constituents in Republican districts. Given this, here we examine whether the changes in the voting behavior of incumbent Republicans we reported in Table 6 of

the main text can still be rationalized as a product of trade-induced changes in the policy preferences of their constituents.

We do so by estimating a series of regressions analogous to our preferred specifications from Table 4 of the main text, in which we allow the effects of the US and Mexican tariff cuts to vary on the basis of the political party of the district's (or county's) representative. These results are presented in Table E.1. Panel (a) reports our estimates of NAFTA's effects on stated support for spending on environmental protection. Panels (b) and (c) report the corresponding estimates for income per capita, and ambient pollution concentrations, respectively. In each panel, the first column reports estimates for the full sample of data; as a result, these estimates capture the average effects of US and Mexican tariff cuts across Democratic and Republican held districts (or counties, in panels (b) and (c)). The specification reported in the second column restricts the sample to the set of "continuing" districts (or counties) that are held by a single party (or legislator) throughout our period of study, while the third restricts the sample to the set of "non-continuing" districts (or counties) that change parties (or legislators) at least once during our period of study.[14] We include controls corresponding to the analogous preferred specification in Table 4, and standard errors clustered by state are reported in parentheses.

Three key findings emerge from Table E.1. First, reductions in US tariffs caused a decrease in support for environmental policy in Republican represented districts and counties, and this effect is larger in incumbent districts and counties that were represented by the Republican party throughout our period of study. For example, our estimates indicate that a 1 pp reduction in US tariffs led to a 16.2 pp reduction in the likelihood a respondent supported increased spending on environmental protection across all Republican districts, but a 42.3 pp reduction in Republican districts in our continuing sample. Our estimates for income per capita and ambient pollution concentrations exhibit a similar pattern. This suggests that our finding that incumbent Republicans reduced their support for environmental policy in response to reductions in US tariffs can indeed be rationalized as a product of trade-induced changes in constituent preferences.

Second, reductions in US tariffs appear to have had, at most, a limited effect on the demand for environmental policy in our non-continuing sample, regardless of the district's party. For example, the estimates reported in column (3) indicate that a 1 pp reduction in US tariffs led to a 7.5 pp reduction in the likelihood a respondent supported increased spending on environmental protection in Republican districts and a 3.6 pp reduction in Democratic districts, although these estimates are not statistically significant at conventional levels. Our corresponding estimates for income per capita and ambient pollution concentrations are also small when compared with the estimates from our sample of continuing counties. These findings suggest that changes in constituent demands for environmental policy are unlikely to explain the observed shift away from the Democratic party in response to NAFTA, a point to which we return in Section E.2.2.

The third, and final, key finding that emerges from Table E.1 is that reductions in US tariffs also appear to have reduced the demand for environmental policy amongst

[14] As the ANES samples individuals from a subset of districts in each year, we define continuing districts as those represented by a single party, rather than single representative, over our sample.

Table E.1: The Effects of NAFTA on the Demand for Environmental Policy, by Party

| | Panel (a): Support for Env. Prot'n | | | Panel (b): Income Per Capita | | | Panel (c): ln(TSP) | | |
|---|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
| $\Delta\tau_r^{USA}$ x $Post_t$ x Rep. | -0.162 | -0.423 | -0.075 | -91.515 | -319.985 | -57.721 | -0.062 | -0.131 | -0.054 |
| | (0.062) | (0.144) | (0.113) | (39.357) | (90.058) | (41.683) | (0.023) | (0.154) | (0.014) |
| $\Delta\tau_r^{USA}$ x $Post_t$ x Dem. | -0.137 | -0.225 | -0.036 | -88.405 | -190.892 | -41.681 | -0.049 | 0.046 | -0.030 |
| | (0.063) | (0.190) | (0.108) | (39.053) | (47.614) | (39.035) | (0.020) | (0.054) | (0.013) |
| $\Delta\tau_r^{Mex}$ x $Post_t$ x Rep. | -0.030 | 0.052 | 0.030 | 7.420 | 16.946 | 4.191 | 0.009 | 0.031 | 0.007 |
| | (0.028) | (0.064) | (0.060) | (6.385) | (9.030) | (8.187) | (0.004) | (0.028) | (0.005) |
| $\Delta\tau_r^{Mex}$ x $Post_t$ x Dem. | -0.023 | -0.020 | 0.015 | -3.964 | -10.149 | -3.858 | 0.012 | -0.077 | 0.007 |
| | (0.018) | (0.071) | (0.053) | (4.412) | (30.574) | (5.120) | (0.007) | (0.051) | (0.006) |
| Continuing | | X | | | X | | | X | |
| Non-Continuing | | | X | | | X | | | X |
| $R^2$ | 0.20 | 0.24 | 0.25 | 0.95 | 0.97 | 0.95 | 0.89 | 0.86 | 0.91 |
| Obs. | 7766 | 4359 | 3407 | 28649 | 6226 | 22423 | 1291 | 358 | 932 |

*Notes*: Table shows results of the NAFTA tariff reductions on stated views on environmental policy (Panel (a)), local economic conditions (Panel (b)), and local environmental conditions (Panel (c)), allowing the effects to vary by each region's political party. The dependent variable in Panel (a) is an indicator for whether a survey respondent believes the federal government should increase spending on environmental protection. Data is taken from the American National Election Studies (ANES) survey. The dependent variable in Panel (b) is the county's average income per capita. Data is taken from the Bureau of Economic Analysis's Regional Economic Accounts. The dependent variable in Panel (c) is the natural log of the county's median daily ambient total suspended particulate concentration. Data is taken from the Environmental Protection Agency's Air Quality System. In each Panel, the first column shows the results allowing the effect of NAFTA's tariff reductions to vary by the party held by the region's representative. The second column in each panel restricts this analysis to regions represented by the same party (Panel (a)) or legislator (Panel (b) and (c)) throughout. The third column in each panel restricts this analysis to regions that change parties (Panel (a)) or legislators (Panel (b) and (c)). All regressions in Panel (a) include congressional district fixed effects, year fixed effects, voter and interviewer characteristic by year fixed effects, initial district economic, regulatory, and demographic by year fixed effects, an indicator of whether the district is held by democrats, and are weighted by the ANES sample weights. All regressions in Panels (b) and (c) include county and year fixed effects, initial county economic, regulatory, and demographic by year fixed effects, and an indicator of whether the county is in a district held by democrats. Standard errors clustered by state are reported in parentheses.

constituents represented by Democrats. For example, the estimate reported in the second row of column (1) indicates that a 1 pp reduction in US tariffs led to a 13.7 pp reduction in the likelihood a respondent supported increased spending on environmental protection in Democratic districts, while the corresponding estimates in columns (4) and (7) suggest that a 1 pp reduction in US tariffs decreased per capita incomes and ambient pollution levels in affected Democratic counties by close to 88 dollars and 4.9%, respectively. However, recall that the results presented in Table 6 indicate incumbent Democratic legislators, unlike their Republican counterparts, do not change their voting behavior on environmental bills in response to reductions in US tariffs.

To investigate this discrepancy further, we exploit the fact that the ANES also contains information as to whether each respondent is a member of the Democratic or Republican parties, or identifies as an independent. This allows us to examine whether there is heterogeneity in the effects of the NAFTA tariff cuts on stated support for environmental policy across voters with different political affiliations.

To do so, we estimate three regressions analogous to our preferred specification from panel (a) of Table 4, but we now allow the effects of the US and Mexican tariff reductions to vary by both the political party of the district's representative, as well as the respondent's self-reported political affiliation. These results are reported in Table E.2. As in panel (a) of Table E.1, we first examine the full sample (column (1)), and then our continuing and non-continuing samples (columns (2) and (3), respectively). All specifications include controls corresponding to the preferred specification from panel (a) of Table 4, and in all cases, standard errors clustered by state are reported in parentheses.

The estimates reported in Table E.2 indicate that the reductions in the demand for environmental policy in response to US tariff reductions documented in Table E.1 are driven by the responses of constituents who self identify as either a Republican or an Independent. For example, the estimates reported in the first three rows of column (1) suggest that a 1 pp reduction in US tariffs led to large reductions in the likelihood of supporting spending on environmental protection amongst Independents and Republicans (reductions of 23.5 pp and 16.9 pp, respectively). In contrast, a 1 pp reduction in US tariffs only led to a 4.1 pp reduction in the likelihood of support amongst Democrats in these districts, although this effect is not statistically significant at conventional levels. A similar pattern arises in districts represented by Democratic legislators; a 1 pp reduction in US tariffs led to 7 pp, 23.9 pp, and 14.3 pp reductions in the likelihood of supporting spending on environmental protection amongst Democrats, Independents, and Republicans, respectively (although only the estimate for Independents is statistically significant). These results provide a natural explanation for the observed difference in the change in voting of incumbent Republican and Democratic legislators in response to the US tariff cuts: Democratic legislators appear not to change their votes because trade liberalization does not impact the environmental policy demands of their main political constituency.

A remaining concern is that our estimates of NAFTA's effects on voting by incumbent legislators could also be capturing differential concerns over industrial flight across districts represented by Democrats and Republicans, due to factors such as differential lobbying in response to trade. To address this concern, we estimate the effects of NAFTA's tariff reductions on RCVs by incumbent politicians, allowing the effects of

Table E.2: Heterogeneity in NAFTA's Effects on Voters' Views on Environmental Policy

| | (1) | (2) | (3) |
|---|---|---|---|
| $\Delta\tau_r^{USA}$ x $Post_t$ x Rep. Dist. | | | |
| x Dem. Voter | -0.041 | -0.200 | 0.045 |
| | (0.066) | (0.170) | (0.129) |
| x Ind. Voter | -0.235 | -0.633 | -0.135 |
| | (0.081) | (0.177) | (0.124) |
| x Rep. Voter | -0.169 | -0.240 | -0.122 |
| | (0.089) | (0.166) | (0.157) |
| $\Delta\tau_r^{USA}$ x $Post_t$ x Dem. Dist. | | | |
| x Dem. Voter | -0.070 | -0.003 | -0.040 |
| | (0.088) | (0.230) | (0.132) |
| x Ind. Voter | -0.239 | -0.424 | -0.008 |
| | (0.088) | (0.175) | (0.126) |
| x Rep. Voter | -0.143 | -0.354 | -0.013 |
| | (0.093) | (0.165) | (0.149) |
| $\Delta\tau_r^{Mex}$ x $Post_t$ x Rep. Dist. | | | |
| x Dem. Voter | -0.034 | -0.012 | 0.031 |
| | (0.034) | (0.062) | (0.065) |
| x Ind. Voter | -0.020 | 0.079 | 0.030 |
| | (0.027) | (0.065) | (0.060) |
| x Rep. Voter | -0.031 | 0.003 | 0.047 |
| | (0.033) | (0.058) | (0.070) |
| $\Delta\tau_r^{Mex}$ x $Post_t$ x Dem. Dist. | | | |
| x Dem. Voter | -0.043 | -0.096 | -0.005 |
| | (0.024) | (0.083) | (0.053) |
| x Ind. Voter | 0.016 | 0.027 | 0.047 |
| | (0.025) | (0.063) | (0.061) |
| x Rep. Voter | -0.030 | -0.019 | 0.007 |
| | (0.023) | (0.065) | (0.059) |
| $R^2$ | 0.22 | 0.25 | 0.26 |
| Obs. | 7766 | 4359 | 3407 |

*Notes*: Table reports estimates of the effects of NAFTA tariff reductions on views expressed on environmental policy stringency. Panel (a) reports estimates using all districts, Panel (b) reports estimates using the sub-sample of districts that are always represented by the same party over our period of study, and Panel (c) reports estimates using the sub-sample of districts whose party changes over our period of study. The dependent variable in all regressions is an indicator of whether the survey respondent believes the federal government should increase spending on environmental protection. All regressions include congressional district and year fixed effects, baseline district Clean Air Act non-attainment status and characteristic trends, voter and interviewer demographic trends, an indicator of whether the district is held by democrats, controls for the voter's party affiliation, and are weighted using the ANES sample weights. Standard errors clustered by state are reported in parentheses.

treatment to vary by both the representative's party and their district's specialization in polluting industries. To do so, we estimate regressions based on Equation (1), allowing the effect of treatment to vary by the representative's political party and whether the initial average cost of complying with environmental regulation for industries in the district is relatively high or low. We also include legislator fixed-effects to restrict our analysis to changing views among incumbent politicians.

The results of this exercise are displayed in Table E.3. As in Table 3 of the main text, each of the four columns corresponds to a different measure of the costs of regulatory compliance. In Panel (a), our measure is the ratio of pollution abatement operating costs (PAC) to the total cost of materials. In Panel (b), our measure is the share of PAC in industry value added. In each panel, the first column classifies districts as having relatively high or low compliance costs (High and Low PAC, respectively) if the average costs of complying with environmental regulation are above or below that of the average district; the second column makes this delineation based on the cost of regulatory compliance for the median district.

These estimates indicate that the responses by incumbent politicians did not reflect concerns over industrial flight. We find that, for both parties and across each of our four specifications, the estimated effects of the reduction in US import tariffs following NAFTA for Low and High PAC districts are not statistically distinguishable from each other. Similarly, we also find no significant differences across Low and High PAC districts for the Mexican tariff changes.

### E.2.2 NAFTA and Changes in Partisan Representation: Redux

The estimates reported in Table E.2 also provide further evidence that that trade-induced changes in partisan representation are unlikely to be due to the effects of tariff reductions on constituent preferences for environmental policy, as both US and Mexican tariff cuts have little effect on voter views in non-continuing districts. Hence, we next investigate whether the change in partisan representation, that resulted in a change in voting on environmental RCVs, is consistent with affected voters reducing support for Democratic legislators for adopting pro-NAFTA positions prior to the agreement's ratification.

There is reason to believe that this type of response could underlie the change in partisan representation. As we noted in the main text, the work of Choi et al. (2022) suggests that voters in regions most affected by NAFTA were more likely to switch from supporting Democrats to Republicans due to the former party's support for the agreement. Given the stark difference in support for environmental issues across the two parties, this decrease in support for Democrats could manifest as a reduction in pro-environmental RCVs if it led to the election of more Republicans.

We investigate this possibility by again estimating the effects of the NAFTA tariff cuts on electoral outcomes using Equation (5), and our data on electoral results from the MIT Election Data Lab. However, we now consider the effects of the tariff cuts across two sub-samples differentiated according to whether the district's representative voted for or against the NAFTA Implementation Act (HR 3450), the roll call vote to ratify the agreement. In the first, we restrict the sample to the set of districts whose representative opposed NAFTA, while in the second, we restrict the sample to the set of districts whose

Table E.3: The Effects of NAFTA on Roll Call Votes by Party and Abatement Cost

| | Panel (a): PAC/Materials Costs | | Panel (b): PAC/Value Added | |
|---|---|---|---|---|
| | (1) | (2) | (3) | (4) |
| $\Delta\tau_r^{USA}$ x $Post_t$ x Rep. | | | | |
| x High PAC | -0.128 | -0.090 | -0.116 | -0.070 |
| | (0.066) | (0.067) | (0.064) | (0.041) |
| x Low PAC | -0.094 | -0.112 | -0.111 | -0.121 |
| | (0.049) | (0.057) | (0.050) | (0.061) |
| $\Delta\tau_r^{USA}$ x $Post_t$ x Dem. | | | | |
| x High PAC | -0.131 | -0.044 | -0.027 | -0.060 |
| | (0.069) | (0.053) | (0.054) | (0.051) |
| x Low PAC | 0.001 | -0.017 | -0.021 | -0.014 |
| | (0.048) | (0.050) | (0.049) | (0.047) |
| $\Delta\tau_r^{Mex}$ x $Post_t$ x Rep. | | | | |
| x High PAC | -0.008 | -0.013 | -0.016 | -0.020 |
| | (0.023) | (0.024) | (0.020) | (0.022) |
| x Low PAC | -0.009 | -0.013 | -0.012 | -0.010 |
| | (0.010) | (0.011) | (0.013) | (0.012) |
| $\Delta\tau_r^{Mex}$ x $Post_t$ x Dem. | | | | |
| x High PAC | 0.047 | 0.018 | 0.003 | 0.020 |
| | (0.034) | (0.028) | (0.023) | (0.024) |
| x Low PAC | 0.016 | 0.021 | 0.026 | 0.019 |
| | (0.022) | (0.022) | (0.026) | (0.024) |
| $R^2$ | 0.45 | 0.45 | 0.45 | 0.45 |
| Obs. | 50321 | 50321 | 50321 | 50321 |

*Notes*: Table reports estimates of the effects of the NAFTA tariff reductions on roll call votes in the House of Representatives allowing the effects to vary by party and across districts on the basis of their average costs of complying with environmental regulation. The dependent variable in all regressions is an indicator of whether the roll call vote cast by a representative on a particular bill is pro-environment. In Panels (a) and (b), the cost of complying with environmental regulation are measured as the ratio of PAC to the total cost of materials, and the ratio of PAC to value added, respectively. In the first column of each panel, districts are classified as having relatively high or low compliance costs (High and Low PAC, respectively) if the average costs of complying with environmental regulation are above or below that of the average district. In the second column of each panel, districts are classified as having High or Low PAC if the average costs of complying with environmental regulation are above or below that of the median district. In each column, the first two rows show the effect of a reduction in US import tariffs for districts currently with a Republican or Democratic representative, respectively.The last two rows show the effect of a reduction in Mexican import tariffs for districts currently with a Republican or Democratic representative, respectively. All regressions include district, representative, and year fixed effects, as well as controls for the effects of the Clean Air Act and differential trends in baseline characteristics. Standard errors two-way clustered by state and bill are shown in parentheses.

representative supported NAFTA. For each sample, we examine whether the NAFTA tariff cuts affected the likelihood that the district flipped from Republican to Democrat, and from Democrat to Republican.

The results from this exercise are reported in the two panels of Table E.4. Panel (a) reports estimates of the effects of the NAFTA tariff cuts on the likelihood a district flipped from Republican to Democrat, while panel (b) reports estimates of these effects

Table E.4: The Effects of NAFTA on Electoral Outcomes, by NAFTA Vote Status

| | Panel (a): Pr(Change Rep. to Dem.) | | Panel (b): Pr(Change Dem. to Rep.) | |
|---|---|---|---|---|
| | (1) | (2) | (3) | (4) |
| $\Delta\tau_r^{USA}$ x $\text{Post}_t$ | 0.071 | -0.028 | 0.028 | 0.190[a] |
| | (0.060) | (0.045) | (0.050) | (0.068) |
| $\Delta\tau_r^{Mex}$ x $\text{Post}_t$ | 0.007 | 0.003 | -0.015 | 0.019 |
| | (0.021) | (0.008) | (0.021) | (0.016) |
| Pro-NAFTA Vote | | X | | X |
| $R^2$ | 0.25 | 0.24 | 0.24 | 0.23 |
| Obs. | 985 | 1159 | 985 | 1159 |

*Notes*: Table shows results of the NAFTA tariff reductions on election outcomes in the House of Representatives for the 102nd to the 106th congress, splitting the sample by the representative's vote status on the NAFTA Implementation Act (HR 3450). The dependent variable in Panel (a) is an indicator for whether the district changed from the Republican to Democratic party in the last election. The dependent variable in Panel (b) is an indicator for whether the district changed from the Democratic to Republican party in the last election. The first column in each panel restricts the sample to districts whose representative opposed NAFTA, and the second column in each panel restricts the sample to districts whose representative voted in favor of NAFTA. All regressions include district and year fixed effects and district baseline characteristic and CAA non-attainment trends. All regressions are restricted whose representative voted on HR 3450. Standard errors clustered by state are shown in parentheses.

for districts that flipped from Democrat to Republican. In all cases, standard errors clustered by state are reported in parentheses.

Consistent with Che et al. (2022), the results presented in Table E.4 suggest that NAFTA caused voters to reduce support Democratic representatives who voted in favor of NAFTA. For example, the results in column (4) show that among the sample of districts whose representative voted in favor of NAFTA, a 1 pp reduction in import tariffs caused a 19 pp increase in the likelihood of the district flipping from Democrats to Republicans. In contrast, the results in column (3) show no significant change in party among the anti-NAFTA districts. Moreover, a Wald test comparing these two coefficients indicates that these differences are statistically significant (p-value = 0.07). In addition, the table suggests no punishment occurred for Republicans, with the results in Panel (a) showing no significant effect of either tariff change on the likelihood of a district switching from the Republican to Democratic party in either sub-sample. Together, these results suggest that much of NAFTA's effect on the formation of environmental policy in the US House of Representatives is an incidental byproduct of voters electing Republicans to replace pro-NAFTA Democrats.

## E.3 Voter Views on Reproductive Rights

As we discuss in Section 5 of the main text, our findings indicate that incumbent Democrats may be less likely to cast a pro-choice RCV in response to US tariffs, and we note that this appears to be due to these Democrats responding to changes in the preferences of self-identified Independent voters in their districts. The corresponding results that suggest this are displayed in Table E.5, which reports estimates of the effects of the NAFTA tariff reductions on expressed views on reproductive rights, using the public opinion

data from the ANES. In the table, Panel (a) reports estimates using all districts, while Panel (b) reports estimates using the sub-sample of districts that are always represented by the same party over our period of study, and Panel (c) reports estimates using the sub-sample of districts whose party changes over our period of study. In all cases, the dependent variable is an indicator of whether the survey respondent believes the federal government should allow abortion. All regressions include congressional district and year fixed effects, baseline district Clean Air Act non-attainment status and characteristic by year fixed effects, and voter and interviewer demographic by year fixed effects, and are weighted using the ANES sample weights. In the table, standard errors clustered by state are reported in parentheses.

As the estimates reported in column (1) of Table E.5 show, the reduction in US tariffs only had a meaningful effect on the views of Independent voters that reside in districts held by Democrats. While imprecisely estimated, the results reported in columns (2) and (3) suggest that this effect is likely driven by independent voters in districts that are held by Democrats throughout our period of study, suggesting that the responses of incumbent Democratic legislators that we observe in Table 8 of the main text may be due to these legislators responding to the demands of an important electoral constituency on this issue.

Table E.5: Heterogeneity in NAFTA's Effects on Voters' Views on Reproductive Rights

| | (1) | (2) | (3) |
|---|---|---|---|
| $\Delta\tau_r^{USA}$ x $\text{Post}_t$ x Rep. Dist. | | | |
| x Dem. Voter | -0.000 | 0.209 | -0.179 |
| | (0.084) | (0.181) | (0.154) |
| x Ind. Voter | 0.014 | 0.182 | -0.126 |
| | (0.106) | (0.195) | (0.155) |
| x Rep. Voter | 0.016 | -0.069 | -0.024 |
| | (0.083) | (0.167) | (0.181) |
| $\Delta\tau_r^{USA}$ x $\text{Post}_t$ x Dem. Dist. | | | |
| x Dem. Voter | 0.065 | 0.257 | -0.139 |
| | (0.082) | (0.140) | (0.172) |
| x Ind. Voter | -0.124 | -0.116 | -0.034 |
| | (0.079) | (0.098) | (0.136) |
| x Rep. Voter | 0.020 | 0.054 | 0.163 |
| | (0.085) | (0.137) | (0.157) |
| $\Delta\tau_r^{Mex}$ x $\text{Post}_t$ x Rep. Dist. | | | |
| x Dem. Voter | 0.059 | 0.074 | -0.037 |
| | (0.031) | (0.066) | (0.086) |
| x Ind. Voter | 0.044 | 0.056 | -0.051 |
| | (0.030) | (0.070) | (0.080) |
| x Rep. Voter | 0.029 | 0.039 | -0.041 |
| | (0.022) | (0.061) | (0.073) |
| $\Delta\tau_r^{Mex}$ x $\text{Post}_t$ x Dem. Dist. | | | |
| x Dem. Voter | 0.036 | 0.014 | 0.039 |
| | (0.026) | (0.054) | (0.037) |
| x Ind. Voter | 0.032 | 0.048 | 0.000 |
| | (0.023) | (0.045) | (0.055) |
| x Rep. Voter | 0.031 | 0.024 | 0.015 |
| | (0.025) | (0.054) | (0.056) |
| $R^2$ | 0.23 | 0.27 | 0.28 |
| Obs. | 8761 | 4871 | 3890 |

*Notes*: Table reports estimates of the effects of NAFTA tariff reductions on views expressed on reproductive rights. Panel (a) reports estimates using all districts, Panel (b) reports estimates using the sub-sample of districts that are always represented by the same party over our period of study, and Panel (c) reports estimates using the sub-sample of districts whose party changes over our period of study. The dependent variable in all regressions is an indicator of whether the survey respondent believes the federal government should allow abortion. All regressions include congressional district and year fixed effects, baseline district Clean Air Act non-attainment status and characteristic trends, voter and interviewer demographic trends, an indicator of whether the district is held by democrats, and controls for the voter's party affiliation, and are weighted using the ANES sample weights. Standard errors clustered by state are reported in parentheses.

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